THE REAL EXCHANGE IN THE SHORT, MEDIUM AND LONG RUN

by

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Abstract: This paper uses long-horizon autocorrelations and variance ratio statistics to test for long-term mean reversion in real exchange rates. Unlike most previous tests of this hypothesis, the tests do reject a random walk for monthly data in the post-Bretton Woods era; however, the statistics indicate that positively-correlated innovations, rather than mean reversion, are the source of the rejection. Tests using annual data for the twentieth century also reject the random walk. In this case, however, the rejection can be attributed to mean reversion and confirms that PPP is a long-term phenomenon.

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Introduction

Considerable effort has been spent by economists on testing the theoretical concept of Purchasing Power Parity (PPP) and (in its long-term version) the implication that real exchange rates should exhibit mean reversion.[1] A competing theory, ex ante PPP, has been put forth, but has the unappealing implication that the real exchange rate is a martingale so that the sample path and the unconditional variance of the levels of the exchange rate are unbounded and, even in the long run, neither absolute nor relative PPP necessarily holds.

Much of the empirical work on ex ante PPP has concentrated on the serial independence of real exchange rate innovations in the short run. Using monthly observations for the most part, little evidence has emerged that these innovations are correlated leading many to accept the theory. A similar question regarding martingales has also persisted in the literature on stock returns despite theoretical reasons against such behavior.[2] Recently, research on the behavior of stock prices has produced evidence at odds with the random walk hypothesis. Using autocorrelations from long-run stock returns, Fama and French (1987) conclude that the empirical evidence is consistent with a model of stock price behavior which contains a slowly mean-reverting component. Lo and MacKinlay (1988a and 1988b) and Poterba and Summers (1988) use

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[1] Officer (1976) reviews the early literature on PPP. More recent empirical work will be discussed below.

[2] Lucas (1978) develops a general equilibrium asset pricing model in which asset prices do not follow a random walk.

variance ratio tests of the random walk hypothesis[3] and are able to reject the hypothesis of no serial correlation in stock price innovations, but unlike the negative serial correlation in returns that Fama and French found, positive serial correlation is the source of the rejection.

Both of these tests can provide valuable information on the long-term behavior of a time series which other statistics are unable to identify. This paper employs these tests in order to examine the behavior of the real exchange rate from both a short and long-term perspective using monthly and annual data. Results using monthly short-term differences generally do not reject the martingale hypothesis, which agrees with previous work. At differences of from 16 to 32 months, however, statistically significant deviations from the null hypothesis are found in most countries tested. These rejections, however, do not support short-term mean reversion, but rather point to positive serial correlation in the real exchange rate innovations which occurred during the 1980's. At odds with these monthly results, the annual time series also reject the martingale hypothesis, but in this case the evidence is in favor of mean reversion and long-run PPP.

The outline of the paper is as follows. The next section describes the theoretical background and previous empirical research. Section II investigates the autocorrelations of real

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[3] Financial economists tend to use the terms random walk and martingale interchangeably. The definitions used by Karlin and Taylor (1975) differentiate between the two concepts. In what follows the two will be taken to be the same.

exchange rates. Section III describes the variance ratio statistics and discusses the results of applying them to the data. Conclusions are presented in the final section.

II. Background

PPP can be summarized by the relationship

(1) $P_t = kS_tP_t{*}$

where

S_t = the exchange rate at time t, defined as the home currency price of foreign currency

P_t ($P_t{*}$) = price level in the home (foreign) country

and k is equal to one if absolute PPP holds. The relative version of PPP allows k to differ from one in order to allow for distortions such as tariffs. If these distortions are constant then k is also constant and (1) will hold in rate-of-change form. Critics of PPP in either form argue that there are theoretical models which indicate that one need not expect k to be constant over time. Supporters generally accept this, but believe that any deviations are only transitory and that in the long run the relationship is valid, at least in its relative form.

When k is constant, an implication of (1) is that relative PPP must be expected to hold ex ante. That is,

(2) $E_t[d(t,t-i)] \equiv E_t[r_t - r_{t-i}] = 0$

where

$r_t \equiv \ln(S_t) - \ln(P_t) + \ln(P_t{*})$ = real exchange rate at time t

$d(t,t-i)$ is the i period innovation in r_t and $E_t[\cdot]$ is the time t expectation operator. Under ex ante PPP the real exchange rate will be a martingale, which implies that its innovations are serially uncorrelated and not predictable.

Roll (1979) observes that, under certain assumptions, ex ante PPP is a consequence of intertemporal commodity arbitrage. Using monthly observations, Roll (1979), Frenkel (1981) and Cumby and Obstfeld (1984) find no evidence that changes in the real exchange rate are serially correlated. Adler and Lehmann (1983) develop the martingale property on the basis of financial market efficiency and are also unable to detect serial correlation in real exchange rate innovations using either monthly observations during the post-Bretton Woods era or annual observations for the twentieth century.[4]

The inability to detect serial correlation in real exchange rate innovations could attest to the validity of ex ante PPP, or alternatively reflect the low power of the statistical procedures employed.[5] In support of the low-power alternative, Mishkin (1984) rejects joint tests of ex ante PPP and uncovered interest rate parity, the rejection presumably due to the increased power of the joint test. Cumby and Obstfeld (1984) reject ex ante PPP on the basis of a test of the coefficients from a regression of inflation differentials on exchange rate changes, again with the

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[4] Adler and Lehmann use the annual data set compiled by Lee (1976), which consists of average annual exchange rates and price indices. As discussed in more detail below, the averaging process causes statistical problems which makes interpretation of their results difficult.

[5] Hakkio (1986) and Poterba and Summers (1988) present Monte Carlo evidence which suggest that tests for a random walk have low power.

help of a stronger null hypothesis. The findings of Abuaf and Jorion (1990) lend further support to the hypothesis that previous tests had low power since they find evidence of mean reversion in annual observations of the real exchange for the twentieth century using jointly estimated regression equations.

An additional criticism of most previous work is that it has concentrated on testing for linear dependence in exchange rate innovations. As discussed below, general equilibrium models typically produce nonlinear exchange rate specifications so that, theoretically, there is no reason to limit empirical work to tests of linear dependence. Hsieh (1989) looks at nominal daily exchange rates using the statistic developed by Brock, Dechert and Scheinkman (1987) (BDS) to test for general forms of dependence. While Hsieh is able to reject the hypothesis of no dependence, the BDS statistic is sensitive to any form of dependency, linear or nonlinear. Hsieh attempts to rule out linear dependence by filtering the data, but it is still possible that the observed dependence is linear. The BDS statistic also requires a large number of observations, which are unavailable at the monthly or annual frequency and so limit the viability of employing it in tests of PPP.

Much early work on stock market efficiency also concentrated on serial correlation of returns in that market.[6] While the majority of such tests have been unable to uncover any evidence of

[6] See Fama (1970) for a review of the early efficient markets literature.

predictability, recent work has cast new light on the area. In order to test for a slowly mean-reverting process in stock returns, Fama and French (1987) use long-horizon returns in regression equations and find significant negative serial correlation. Interestingly, the Fama and French findings are directly at odds with those of Lo and MacKinlay, whose variance ratio tests indicate positive rather than negative serial correlation. Confirming both of these findings, Poterba and Summers (1988) find that stock returns exhibit positive serial correlation over short horizons and negative correlation over long horizons.

Some empirical work on exchange rates has been conducted using variance ratio tests, but not with the long-term autocorrelations approach. Huizinga (1987) used a spectral estimator of the variance ratio for several currencies against both the U.S. dollar and the British pound. Without well-developed distribution theory for that estimator, however, no statistically significant departures from the random walk hypothesis were found. Kaminsky (1987) employs one version of the Lo and Mackinlay variance ratio, but only for monthly observations of four currencies against the U.S. dollar and three currencies against the pound and without allowing for conditional heteroskedasticity in the innovations. Using the period March 1973 to April 1986, little evidence against the random walk hypothesis is uncovered.

The disturbing implication of ex ante PPP as a theory of real exchange rate movements is that, with no tendency toward mean reversion, acceptance implies that absolute and relative PPP, even in the long run, are rejected. It may be, however, that constancy of k (in (1)) is overly restrictive and movement in either real or nominal variables results in temporary or permanent changes in the value of k. In fact, alternative theories of the real exchange rate exist which suggest that k is not constant.

The "productivity differential" model (see Balassa (1964) and Samuelson (1964)) attributes movements in the real exchange rate to labor productivity differentials in the traded and non-traded goods sectors and the wages of two countries. Similarly, the "middle-products" model of Sanyal and Jones (1982) finds that the value of k is roughly proportional to the ratio of the two countries' wages. Exogenous shocks to productivity, or wages depending on the model, in either country will result in real exchange rate movement. To the extent that the shocks persist, no reversion in the real exchange rate will occur in this framework. Furthermore, unless the shocks are forecastable, neither will innovations in the real exchange rate be. Ex post, however, if these shocks are somehow correlated, then so will movements in the real exchange rate will be. Mean reversion would be expected if capital and/or labor is mobile and can move to offset any differentials in productivity or wages; however, frictions in these markets could make such mean reversion a long-term, rather than short-term phenomenon.

A rational expectations model of real exchange rate behavior which receives much attention is the overshooting model of Dornbusch (1976). In that model, due to sticky prices, monetary shocks will cause the exchange rate to overshoot its long-run equilibrium value, leading to positive serial correlation in the real exchange rate. An alternative early rational expectations model of exchange rate behavior, but with flexible prices and uncertainty, was developed by Mussa (1977, 1982), who found that the real exchange rate is a function of the current value, as well as the expectations of all future values of innovations in aggregate demand, output and the world rate of interest. To the extent that shocks or changes in expectations are serially correlated or exhibit mean-reverting behavior, so to will the real exchange rate exhibit such behavior.

Stockman (1980), Lucas (1982) and Svensson (1985) develop general equilibrium rational expectations models of exchange rates based on cash-in-advance models of money and optimizing behavior on the part of individual agents. In these models the value of the real exchange rate is determined by the expected marginal utilities of the domestic and foreign goods, and the expected domestic and foreign rates of inflation. Thus, both current and expected future price movements and goods shocks will affect the real exchange rate. While the general equilibrium nature of these models is intuitively appealing, they pose problems for empiricists due to the nonlinearities in their solutions and their dependence on unobservable expectations and preferences.[7] Like the earlier rational expectations model of Mussa, serial correlation or mean-reverting behavior would be derived from the processes of the underlying state variables. Since the models are based on endowment economies, capital and labor mobility and the mean reversion that they could induce can not be addressed in these models.

Stulz (1987) extends the rational expectations framework to include nontraded goods and heterogeneous agents. He finds that the real exchange rate is determined by the ratio of the stocks of nontraded goods in the two countries. Martingale behavior will be observed if the means and variances of the growth rates of the stocks of nontraded goods are constant and identical across countries. Otherwise, serial correlation in real exchange rates can occur.

Dumas (1989) examines a model in which deviations from PPP are an equilibrium phenomenon due to the presence of fixed costs for transporting capital between countries. In the short-term, with independent and identically distributed shocks to output, the exchange rate may appear to be a martingale; however, the behavior of agents actually induces serial correlation in the exchange rate leading to mildly mean-reverting behavior. Thus, it is possible in this model that both short-term martingale behavior and long-

[7] Meese and Rose (1989) investigate the implications of the nonlinearities from this type of model on the exchange rate, but do not find evidence that incorporating the nonlinearities improves their ability to explain movements in the exchange rate.

term mean reversion could be observed ex post.

As we have seen, economics offers several alternatives to the theory of ex ante PPP. Serial correlation in shocks to wages, productivity or output can all lead to serial correlation, either positive or negative, in the real exchange rate. The theories, however, say little about either the forecastability of the innovations of these "state" variables or about the time horizon required to offset any shocks that occur. For example, a real wage shock may cause a contemporaneous real exchange rate movement, but ultimately, if labor and capital are mobile, such a shock should be offset and parity returned. Frictions may prevent this reversion from happening quickly, however, and a slowly mean-reverting process for the real exchange rate may be the reason that previous tests, which looked mainly at short-term changes in the real exchange rate, were unable to find any evidence of mean reversion. While this is supportive of ex ante PPP, it is also possible that the empirical tests used have not been sufficiently powerful or have concentrated on too short a horizon to detect whatever correlation exists. The following sections look at long-horizon autocorrelations and variance ratio tests of real exchange rate movements. Collectively they provide substantial evidence in favor of PPP as a long-run phenomenon.

II. Long-Horizon Real Exchange Rate Changes

The proposition that stock prices follow a random walk has gained acceptance by many. Recently, however, Summers (1986) and

Fama and French (1987) have suggested that stock prices may have slowly decaying stationary components. Fama and French show that a slowly mean-reverting component will induce negative serial correlation in returns and test this hypothesis by considering the autocorrelations of stock returns for increasing holding period returns. They conclude that significant negative serial correlation exists, with the correlation weak for short-term returns, but stronger for long-horizon returns.

Some of the theoretical models of real exchange rate behavior discussed in the previous section also provide for the possibility that short-term real exchange rate movements may exhibit martingale behavior, but that long-term mean reversion will occur. This section conducts tests for a slowly mean-reverting component in real exchange rates by looking at the autocorrelations of long-horizon real exchange rate innovations.

Let $\rho(i)$ be the first-order autocorrelation of the real exchange rate change over an interval of length i, d(t,t-i). Looking at estimates of $\rho(i)$ for increasing values of i may provide evidence on the existence of a slowly mean-reverting process in the real exchange rate. Estimates of $\rho(i)$ are obtained from standard OLS regresion coefficients. In order to obtain as much precision as possible, overlapping differences are employed in the regressions and Hansen and Hodrick (1980) standard errors are used for conducting inferences. Additional adjustments are also needed to account for the bias in OLS estimates as described by Fama and French. These adjustments are obtained from simulations and the reported t-statistics test for deviations from the estimated bias.[8]

The Data

Real exchange rates were generated by obtaining end-of-month nominal exchange rates for nine currencies against the U.S dollar from the International Financial Statistics (IFS) data base. Deflating these by the consumer price indices (CPI) and wholesale price indices (WPI), also from the IFS, produced the real exchange rates.[9] Much debate has occurred over what index is the appropriate one for testing PPP. Keynes (1930) favored the WPI because of its higher proportion of widely-traded commodities compared to the CPI, which has more non-traded services. An alternative view takes exactly the opposite stance and views the CPI as a more reasonable index of purchasing power. For comparison, both will be employed with only marginal differences in the results. In the tests, the natural logarithm of the real exchange rate was used. The monthly time series used covers the period 1973:06 - 1988:12.

Graphs of monthly observations of the real exchange rates for selected countries are presented in Figures 1 through 4. Figures 1 and 2 show the value of the U.S. dollar against the Canadian dollar and German DM respectively. The Canadian dollar has

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[8] Simulations throughout the paper were conducted using the IMSL routine GGNML to generate random normal variables. From these, autocorrelations were calculated. This process was repeated 10,000 times. The average value of the estimated autocorrelation is reported in Tables 1A and 1C as the bias.

[9] WPI values were unavailable for France.

noticeably less variation than the DM and appears to display less serial correlation. Notable in Figure 2 is the rapid rise and subsequent fall in the value of the dollar over the period 1980-87. Except for the Canadian dollar, this rise and fall characterized the behavior of most of the dollar exchange rates in the sample period. For comparison, Figures 3 and 4 present the French franc/German mark and Japanese yen/Swiss franc exchange rates. Evident in the FF/DM graph is the decrease in volatility that occurred subsequent to the introduction of the European Monetary System in 1979, as well as the discrete jumps that periodic realignments within that system have caused.

The first six autocorrelations of the first difference of the CPI real exchange rate is reported in Table 1A. Given the number of observations, very few of these are statistically significant. Moreover, no clear pattern of positive or negative autocorrelations is obvious, except perhaps for the negative correlation at lag 6. More positive values than negative values are observed, but, due to the insignificance of the individual entries, it is not surprising that Adler and Lehmann (1983) are unable to reject the hypothesis that all of the autocorrelations up to lag 18 are jointly zero. The autocorrelations of the first differences of the WPI exchange rates in Table 1B display similar properties.

The annual data collected by Lee (1976) will also be examined.[10] The data consists of annual average real exchange rates for the period 1900-87. Autocorrelations of the first differences appear in Table 1C. Important in analyzing this data is recognition of the implications that the averaging process has on the time series. Working (1960) shows that by averaging a random walk process, positive first-order serial correlation is introduced into the first-difference of the averaged series. For the case of annual averages of monthly observations, this implies a theoretical first-order serial correlation of the first difference of 0.25 if the process was originally a random walk. Deviations from the random walk hypothesis must, therefore, be measured against this benchmark.[11] Consequently, all the reported first-order correlation coefficients exhibit evidence of negative serial correlation, with the French franc, Italian lire and Dutch guilder statistically significant.[12] Beyond the first-order correlation coefficient substantially more of the estimated coefficients are negative than was the case with the monthly data.

Empirical Results

The first-order autocorrelations for various horizons of

[10] The author would like to thank Phillipe Jorion for supplying this data and also for updating the data from 1976 to 1987.

[11] The effects of this serial correlation should also be evident in the regression tests conducted by Adler and Lehmann (1983). Their failure to incorporate this into their statistics seriously reduces the possibility of rejecting the hypothesis of a random walk.

[12] This inference depends on the standard assumption of iid innovations so that the standard error of the autocorrelations is $1/\sqrt{T}$. Given the nature of the data, the iid assumption is questionable.

changes in the CPI real exchange rate are presented in Table 2A together with t-statistics for the hypothesis that the estimated coefficient differs from the simulated bias, which is reported at the bottom of the table. Few entries are significantly different from either the bias or zero. The exceptions to this are long-horizons for the Belgian and French francs, the DM and the guilder, all of which exhibit positive serial correlation at lags 12 and 16. This observed behavior is at odds both with the hypothesis of mean reversion, as well as with the behavior documented by Fama and French for stocks. Also different from the Fama and French findings is the lack of any U-shaped pattern in the autocorrelations. In contrast, the yen/SF exhibits significant negative correlation at long lags. The behavior of the WPI real exchange rates in Table 2B is quite similar to that of the CPI rates. Again, all of the significant statistics are greater than the estimated bias, with the exception of the yen/SF.

The annual data in Table 2C, on the other hand, present a far different picture. All but one of the reported t-statistics are negative, indicating negative serial correlation in the data. Again, statistical significance is more apparent for higher-order lags, but in this case five out of seven exchange rates reject the hypothesis of no correlation at lag 8. Similar to the monthly data, no obvious pattern in the coefficients emerges as it did in the stock market data analyzed by Fama and French.

Richardson (1988) questions the findings of Fama and French by noting that individual t-statistics are not independent and that joint tests which incorporate the correlation between the estimated coefficients are more meaningful. Richardson computes the covariance matrix between estimated coefficients under the null hypothesis and suggests the chisquared statistic reported in the last column of the three tables. The importance of the correlation between estimated coefficients is apparent in Tables 2A and 2B, where none of the joint tests reject the hypothesis that all estimated coefficients are zero. For the annual data, however, three strong rejections are obtained for France, Italy and Japan. Given the signs of the estimated coefficients, this provides evidence in favor of long-term mean reversion in the real exchange rate. The next section pursues the matter further with additional tests of long-term behavior.

III The Variance Ratio

Let X_t be a random variable which follows a diffusion process

$$dX_t = \mu dt + \sigma dW(t).$$

If X_t is sampled at discrete intervals, the variance of the increments is linear in the observation interval, i.e. the variance of $X_t - X_{t-2}$ is twice the variance of $X_t - X_{t-1}$. Lo and MacKinlay (1988a) use this to develop a test of the random walk hypothesis. With homoskedastic increments they show that the statistic

$$(3) \quad M_r(q) = \frac{\sigma^2(q)}{\sigma^2(1)} - 1$$

is asymptotically normally distributed, where

$$(4)\quad \sigma^2(q) = \frac{1}{m}\sum_{k=q}^{nq} [X_k - X_{k-q} - q\mu]^2$$

$m \equiv q(nq-q+1)(1-q/nq)$, $nq+1$ is the total number of observations and the usual maximum-likelihood estimator of μ is employed. Inferences may be drawn through the use of the statistic

$$(5)\quad Z1(q) = M_r(q)\left[\frac{2(2q-1)(q-1)}{3q}\right]^{-\frac{1}{2}}\sqrt{nq}$$

which, asymptotically, has a standard normal distribution. Lo and Mackinlay (1988b) find little evidence that the statistic deviates substantially from its asymptotic distribution provided the value of q is not too large relative to nq.[13] Rather than rely on the asymptotic distribution, however, inferences will be based on the empirical distribution of the statistic derived from simulations as described below.

In order to allow for general forms of heteroskedasticity in the variance of the increments of X_t, Lo and MacKinlay propose using the same statistic $M_r(q)$, but altering the estimate of its variance. The proposed statistic is

$$(6)\quad Z2(q) = M_r(q)/\sqrt{\theta}$$

where

[13] Lo and MacKinlay suggest that values of q less than one eighth the value of nq will provide accurate inferences using the asymptotic distribution.

$$\theta(q) = \sum_{j=1}^{q-1} \left[\frac{2(q-j)}{q}\right]^2 \delta(j)$$

$$\delta(j) = \frac{\sum_{k=j+1}^{nq} (X_k - X_{k-1} - \mu)^2 (X_{k-j} - X_{k-j-1} - \mu)}{\sum_{k=1}^{nq} (X_k - X_{k-1} - \mu)^2} .$$

One implication of this form of the variance ratio test is that it is asymptotically equivalent to a linear combination of the first q-1 autocorrelation coefficients of the first difference of X_t,[14] which will be useful when examining the results obtained below.

There is more than one advantage of the variance ratio statistic over other tests of random walk behavior. First, Faust (1989) and Poterba and Summers (1988) show that the statistic is more powerful than other tests, especially against processes which exhibit slowly mean-reverting behavior. Given the inability of previous test to identify short-term mean reversion, as well as the long-term mean reversion evident in some of the autocorrelations of the previous section, this is a compelling reason to consider the statistic. Second, the statistic can be calculated so that it is consistent under general forms of heteroskedasticity. Thus, any rejections that occur can not be attributed solely to inconsistency due to nonstationarity in the variance of the process. Third, by defining the statistic to allow for overlapping observations, the statistic allows for evaluation of long-term serial correlation using a relatively

[14] See Lo and MacKinlay's proof of Theorem 2.

small number of observations. For these reasons, this variance ratio statistic may provide additional evidence on long-run PPP behavior.

Empirical Results

Results of the tests for the monthly data using the CPI are given in Table 3A where the variance ratio, $M_r(q)+1$, as well as both Z1(q) and Z2(q) are presented. The homoskedastic statistics, Z1(q), are presented in parentheses, while the heteroskedasticity-consistent statistics, Z2(q), are presented in brackets. Although both have asymptotic standard normal distributions, simulations were conducted to verify that the small sample properties were well approximated by the asymptotic distribution. These simulations were carried out by generating a sample of standard normal random variables and then calculating the corresponding variance ratio and Z-statistics. This procedure was repeated 10,000 times to generate an empirical distribution of the statistics. The 5% critical values of the empirical distributions for Z1(q) are always less than the asymptotic values when q is less than 16. Conversely, Z2(q) appears to have somewhat larger critical values than predicted by the asymptotic distribution. In both cases, for q greater than or equal to 16 the critical values are larger than that of a standard normal random variable. Poterba and Summers (1988) question the wisdom of using the conventional 5% critical value for tests of this nature.[15]

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[15] Poterba and Summers state that "[in] order to justify using the conventional 5% test, one would have to assign three times as great a cost to Type I as to Type II errors."

Consequently, the tables indicate significance at both the 5% and 10% levels.

At small increments, q = 2,4 and 8 months, very few rejections of the random walk hypothesis are found. At longer increments, however, the advantage of these statistics becomes evident. When the variance ratio is calculated using increments of 16 months, 6 out of nine currencies (vis a vis the dollar) reject the random walk hypothesis at the 5% level. Using increments of 32 months, all currencies, except for the Canadian dollar, reject the null. Generally, the rejections do not depend on which version of the statistic is employed, in fact the two statistics are usually quite similar in value.[16] Unreported results for q=64 are similar to those for q=32.

These results are generally in agreement with the tests conducted by Huizinga (1987), who found that the variance ratio[17]

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[16] Diebold (1988) conducts tests for heteroskedasticity in both nominal and real exchange rates, concentrating on ARCH specifications. In weekly exchange rates substantial evidence of ARCH is found; however, little evidence is found for either monthly observations of the nominal or real exchange rate. Cumby and Obstfeld (1984), however, found substantial evidence of heteroskedasticity in monthly nominal exchange rates. Unfortunately, tests of heteroskedasticity are generally form specific. In this case, however, the similarity of the two test statistics provides evidence that heteroskedasticity is not a problem. Other tests for heteroskedasticity (unreported) were inconclusive.

[17] Huizinga calculates the variance ratio as a weighted sum of the first N autocorrelation coefficients, where estimation is done for various choices of N. Huizinga employs the weights proposed by Newey and West (1987) in calculating the variance, however, he is forced to truncate his estimator so that his estimator differs slightly from the Lo and MacKinlay version and is not adjusted for heteroskedasticity.

for most currencies he studied increased with the number of autocorrelations used, until a maximum was reached after which the ratio declined. The one exception to this was the Canadian dollar, whose behavior differs in this study as well. The important difference between those results and the present ones is that under the null hypothesis of a random walk, the standard errors used by Huizinga in drawing inferences were so large that in no case was the ratio significantly different from zero. In Table 3A using either the asymptotic or the empirical distribution of the statistics, significant deviations from random walk behavior are observed.

One possibility for the (near) uniformity of the results is that they are all dollar exchange rates and may be highly correlated. To provide some insight into this possibility, two additional exchange rates are examined: the yen/Swiss franc and the French franc/Deutsche mark. The behavior of both of these is somewhat different. The FF/DM results most closely resemble those of the dollar, with the variance ratio generally increasing with the value of q. While the ratio is significantly greater than one for q equal to 16, unlike the dollar exchange rates the variance ratio then declines and becomes insignificantly different from one for q equal to 32. The yen/SF variance ratio also rises with q, for q less than 16, after which it begins to decline and drops below one for q=32. At no point is it significantly in excess of unity. Thus, behavior of the dollar exchange rates is statistically quite different from the two non-dollar exchange rates presented.

For large values of q, the variance ratios in Table 3A generally have values in excess of 1. Lo and MacKinlay show that the variance ratio is asymptotically equal to 1 plus a weighted average of the first q-1 autocorrelation coefficients of $X_t - X_{t-1}$. Thus, for q=2 the variance ratio gives us an estimate of the first-order autocorrelation coefficient. Mean reversion requires negative serial correlation in changes in the exchange rate so that the variance ratio for q=2 should be less than one. Examining Table 3A shows that only 5 out 11 countries have variance ratios less than 1, although none are significantly so. The autocorrelation coefficients in Table 1A confirm this with 6 out of 11 countries producing (insignificantly) negative first-order autocorrelation coefficients. The increase in value of the variance ratio as q increases comes about due to positive higher-order autocorrelation coefficients. This is also evident in Table 1A where nearly two thirds of the autocorrelations are positive. Thus, for periods of up to 32 months there is enough positive correlation in real exchange rate movements that random walk behavior is rejected. This may reflect the swings in the value of the dollar that have occurred since the collapse of Bretton Woods, a good example of which is the general upward trend in the dollar from early 1980 until early 1985, which was then followed by over two years of decline.

Table 3B presents the variance ratios calculated using the WPI. Some difference is noted between the behavior of the two real exchange rates; ultimately, however, when q is equal to 16 five out of the eight dollar exchange rates reject the null hypothesis. Again, the rejections are all due to positive serial correlation in real exchange rate innovations.

The period covered by the monthly data can be conveniently separated into two subperiods: the period prior to the European Monetary System (EMS) and its exchange rate mechanism (1974-1978), and the EMS period during which most of the European countries in the sample actively managed their currencies (against each other). Statistically, it makes sense to look at these two subperiods because of the possible differences in the underlying distributions of the exchange rates under the two regimes. Furthermore, the latter subperiod also experienced remarkably different behavior on the part of the dollar and this behavior may be driving some of the previously discussed results. Unfortunately, due to the small number of observations in the pre-EMS period, only the EMS subperiod can be meaningfully analyzed.

Table 3C presents test results for the EMS subperiod for the CPI and WPI real exchange rates. Due to the lower number of observations, tests are not conducted for q equal to 32. Even with the reduced number of observations, strong rejections of the null hypothesis occur for both real exchange rates. The results show more rejections for EMS countries and CPI real exchange rates and lend support to the hypothesis that the strong appreciation and depreciation of the dollar in the 1980s is responsible for the rejection.

Table 3D presents the variance ratio tests using the annual time series. The skewness in the empirical distribution of the statistic, especially for q=2, confirms that, positive serial correlation is introduced through the averaging process. The results here are surprisingly different from those of the monthly data. First, in only a few instances is the variance ratio greater than unity and in these instances it is not significantly so. Second, while four out of seven countries reject the null hypothesis for q=2 at the 5% level, for larger values of q all countries except Germany reject a random walk. Finally, the difference between the two statistics, Z1 and Z2, is much larger for the annual data than for the monthly data, which is not so surprising given that the annual data cover such different regimes as floating exchange rates, Bretton Woods, the gold standard and two world wars.[18]

The pattern of rejections for the annual data is compatible with theories that the real exchange rate is the sum of a random walk and a stationary mean-reverting process. The implications of this is that innovations are negatively serially correlated for all holding periods and that, up to a certain holding period, the negative serial correlation becomes more negative as the holding period increases. In that case, the variance ratio for q=2 should be less than one and the rejection should be stronger as q

[18] It would be interesting to separate the annual data into subperiods corresponding to each of these regimes; however, this would result in tests involving a very low number of observations, which it seems best to avoid.

increases (up to a certain point). This is similar to the type of behavior observed for the annual exchange rates.[19] It also agrees generally with the autocorrelations reported in Table 1C where only two of the t-statistics are positive. Such behavior is compatible with theoretical models in which short-term deviations caused by exogenous shocks are ultimately followed by mean reversion as differences between the economies are eliminated. An example of this is the model of Dumas (1989), where exogenous shocks can cause short-term stockpiling of goods to occur and short-term martingale behavior, but where those stockpiles will ultimately be consumed (or shipped) and the ratio of the prices of the good in the two countries, i.e. the real exchange rate, will return to unity over time.

IV. Conclusion

Previous tests of the random walk hypothesis for real exchange rates have been unconvincing for proponents of PPP. Unfortunately, the tests involving only short-term approaches appear not to be sufficiently powerful to overcome the problems inherent in tests using near unit-root time series. Using long-horizon autocorrelations and variance ratio tests, however, this paper finds significant deviations from random walk behavior in real exchange rates measured on both a monthly and an annual

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[19] Since the averaging process in the annual data induces spurious first-order serial correlation of about 0.25 into the process, the true autocorrelation coefficient is approximately equal to the variance ratio for $q=2$ less 1.25.

basis. Over the short run these deviations do not provide evidence in favor of mean reversion in real exchange rates, at least not for periods of up to 32 months during the post-Bretton Woods era. In fact, the evidence is in favor of medium-term trends in the real exchange rate traced largely to the behavior of the dollar vis-a-vis the EMS currencies during the 1980s. Tests employing annual data produce strikingly different results. First, the random walk hypothesis is rejected for most countries at lags greater than two years; however, unlike the monthly time series, the annual data exhibits mean reversion.

For those who believe that the random walk hypothesis is a bad theory of exchange rate behavior this paper provides some hopeful evidence. Despite rejections of random walk behavior, however, the tests provide no clear indication of what induces the observed behavior. Why, for example does the Canadian dollar behave so much differently from the other currencies in the short run, but over the longer run displays mean reversion similar to the others? In the theoretical models it could be attributed to similar exogenous shocks for the U.S. and Canada during recent years or, in the Dumas (1989) framework, to lower costs to shipping goods/capital between these physically and economically close neighbors.

On a somewhat more basic level, what theoretical model of real exchange rates would lead to medium-term positive serial correlation followed by long-term mean reversion? It seems unlikely that the standard model of overshooting (Dornbusch

(1976)) could produce these results since in that model a single shock would produce an immediate jump in the exchange rate followed by longer-term reversion. In the 1980's, it seems unlikely that the behavior of the dollar can be attributed to a single exogenous shock, especially since the appreciation that occurred took place over a period of five years. The other models cited above, however, could lead to this sort of ex post behavior given appropriate shocks to the exogenous state variables. Unfortunately, these shocks are unobservable and the models in which they occur, at least those based on optimizing behavior on the part of agents, are generally nonlinear and would require substantial additional assumptions in order to test them empirically. This is clearly beyond the scope of the present analysis, but these findings point to the need to establish more closely the links between real exchange rate changes and innovations in other variables.

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Table 1A

Autocorrelations - (CPI) Monthly Data 1973:06-88:12

lag	1	2	3	4	5	6
BELGIUM	-0.01	0.17*	0.00	0.03	0.04	-0.06
CANADA	-0.08	-0.10	0.07	0.01	0.10	-0.04
FRANCE	-0.06	0.14	0.03	0.09	0.05	-0.05
GERMANY	-0.02	0.14	-0.01	0.02	0.04	-0.07
ITALY	0.02	0.09	0.06	0.03	0.11	-0.07
JAPAN	0.10	0.02	0.10	0.06	0.06	-0.05
NETHERLANDS	-0.02	0.18*	-0.03	0.01	0.05	-0.05
SWITZERLAND	0.01	0.10	0.01	0.01	0.04	-0.08
UK	0.08	0.02	-0.07	0.05	0.04	-0.04
FRANCE/GERM	0.07	0.14	-0.00	-0.11	-0.12	-0.09
JAPAN/SWIT	-0.02	0.01	0.11	0.05	0.04	-0.14

* - significant at the 5% level
The standard errors are $1/\sqrt{181}$ under the nul of iid innovations

Table 1B

Autocorrelations - (WPI) Monthly data 1973:06-88:12

lag	1	2	3	4	5	6
BELGIUM	-0.08	0.16*	-0.08	-0.09	0.01	-0.13
CANADA	-0.09	-0.15*	0.03	0.02	0.09	-0.08
GERMANY	0.00	0.13	-0.05	-0.01	0.04	-0.09
ITALY	-0.01	0.07	0.03	-0.03	0.10	-0.07
JAPAN	0.01	-0.03	0.07	0.06	0.01	-0.10
NETHERLANDS	-0.02	0.17*	-0.06	-0.01	0.03	-0.07
SWITZERLAND	0.01	0.08	-0.02	-0.04	0.03	-0.08
UK	0.10	0.04	-0.08	0.01	0.02	-0.08
JAPAN/SWISS	-0.07	-0.03	0.08	0.06	0.03	-0.16*

* - significant at the 5% level
The standard errors are $1/\sqrt{181}$ under the nul of iid innovations

Table 1C

Autocorrelations - Annual Data 1900-87

lag	1	2	3	4	5	6
CANADA	0.04	-0.32*	0.12	0.10	-0.22*	-0.09
FRANCE	-0.17*	0.09	-0.14	-0.08	-0.23*	-0.01
GERMANY	0.08	-0.01	0.16	-0.15	-0.20	-0.09
ITALY	-0.09*	-0.10	-0.10	-0.08	0.01	-0.07
JAPAN	0.10	-0.20	-0.19	-0.19	0.02	-0.04
NETHERLANDS	-0.03*	-0.07	0.13	-0.26*	-0.12	-0.00
UK	0.13	-0.23*	0.01	-0.10	-0.24*	-0.16

* - significant at the 5% level
The standard errors are $1/\sqrt{81}$ under the null of iid innovations

Table 2A - First-Order Autocorrelations: Monthly Data (CPI) - 1973:06-88:12

	$\rho(1)$	$\rho(2)$	$\rho(3)$	$\rho(4)$	$\rho(6)$	$\rho(8)$	$\rho(12)$	$\rho(16)$	$\chi^2(16)$
	-0.00 (0.00)	0.15 (2.09)**	0.11 (1.19)	0.09 (0.84)	0.17 (1.53)	0.32 (2.94)**	0.38 (3.36)**	0.34 (3.49)**	18.31 (0.31)
	-0.08 (-0.94)	-0.11 (-1.09)	0.03 (0.48)	0.05 (0.77)	0.17 (1.92)*	0.23 (1.74)*	0.18 (0.96)	0.13 (0.84)	22.54 (0.13)
	-0.03 (-0.35)	0.14 (1.84)*	0.19 (1.81)*	0.20 (1.63)	0.14 (1.45)	0.19 (2.10)**	0.25 (2.10)**	0.22 (1.98)**	13.05 (0.67)
	-0.02 (-0.15)	0.13 (1.81)*	0.08 (0.93)	0.07 (0.68)	0.14 (1.10)	0.25 (1.93)*	0.26 (2.38)**	0.18 (2.25)**	18.28 (0.31)
	0.03 (0.38)	0.15 (1.49)	0.16 (1.51)	0.17 (1.42)	0.17 (1.33)	0.20 (1.71)*	0.14 (1.36)	0.04 (1.09)	10.73 (0.83)
	0.08 (1.10)	0.12 (1.63)	0.18 (1.77)*	0.17 (1.37)	0.14 (1.04)	0.13 (1.26)	-0.02 (0.46)	-0.06 (0.30)	9.57 (0.89)
S	-0.02 (-0.19)	0.14 (1.72)*	0.06 (0.68)	0.07 (0.65)	0.13 (1.13)	0.21 (1.79)*	0.24 (2.18)**	0.20 (2.31)**	15.31 (0.50)
D	0.01 (0.15)	0.12 (1.53)	0.10 (1.18)	0.08 (0.90)	0.04 (0.55)	0.09 (0.86)	0.14 (1.37)	0.06 (0.98)	9.76 (0.88)
	0.08 (1.20)	0.03 (0.49)	0.02 (0.33)	0.04 (0.63)	0.07 (0.92)	0.14 (1.43)	0.12 (0.99)	0.14 (1.00)	19.06 (0.27)
	0.08 (1.00)	0.15 (1.33)	-0.01 (0.10)	-0.13 (-0.67)	0.03 (0.43)	0.11 (1.15)	-0.09 (0.05)	-0.31 (-1.67)*	24.81 (0.07)
	-0.03 (-0.33)	0.05 (0.77)	0.16 (1.58)	0.12 (1.12)	-0.04 (0.04)	-0.17 (-1.01)	-0.41 (-2.96)**	-0.55 (-3.28)**	12.47 (0.71)
	-0.005	-0.014	-0.021	-0.029	-0.045	-0.061	-0.092	-0.124	

e 1A for a description of the statistics

Table 2B - First-Order Autocorrelations: Monthly Data (WPI) - 1973:06-88:12

	$\rho(1)$	$\rho(2)$	$\rho(3)$	$\rho(4)$	$\rho(6)$	$\rho(8)$	$\rho(12)$	$\rho(16)$	$\chi^2(16)$
M	-0.10 (-1.21)	0.06 (1.13)	-0.07 (-0.63)	-0.15 (-1.02)	-0.03 (0.08)	0.24 (2.47)**	0.35 (3.43)**	0.29 (3.13)**	29.16 (0.02)
	-0.11 (-1.37)	-0.20 (-2.23)**	-0.05 (-0.30)	-0.03 (-0.01)	0.02 (0.50)	0.09 (0.77)	0.10 (0.65)	0.08 (0.66)	21.58 (0.16)
Y	-0.02 (-0.16)	0.10 (1.40)	0.03 (0.43)	0.01 (0.26)	0.08 (0.65)	0.19 (1.40)	0.22 (2.41)**	0.14 (2.25)**	18.51 (0.29)
	-0.03 (-0.34)	0.10 (1.22)	0.08 (0.81)	0.08 (0.73)	0.12 (0.89)	0.21 (1.69)*	0.22 (2.23)**	0.11 (1.88)*	11.89 (0.75)
	-0.02 (-0.22)	0.03 (0.55)	0.11 (1.30)	0.07 (0.72)	0.00 (0.27)	0.03 (0.62)	-0.04 (0.31)	-0.05 (0.41)	12.85 (0.68)
LDS	-0.03 (-.33)	0.11 (1.58)	0.03 (0.41)	0.02 (0.29)	0.10 (0.75)	0.19 (1.40)	0.23 (2.26)**	0.18 (3.07)**	14.80 (0.54)
LD	-0.02 (-0.21)	0.07 (1.09)	0.02 (0.40)	-0.01 (0.16)	-0.02 (0.12)	0.04 (0.50)	0.13 (1.38)	0.08 (1.24)	9.80 (0.88)
	0.09 (1.38)	0.04 (0.60)	-0.01 (0.09)	-0.02 (0.07)	0.01 (0.45)	0.11 (1.35)	0.10 (1.00)	0.10 (0.95)	17.93 (0.33)
T	-0.08 (-.90)	-0.02 (-0.04)	0.09 (1.05)	0.05 (0.65)	-0.06 (-0.11)	-0.18 (-0.82)	-0.37 (-4.49)**	-0.50 (-2.96)**	16.66 (0.41)

les 1A and 1B for a description of the statistics and the estimated biases.

Table 2C - First-Order Autocorrelations: Annual Data - 1900-87

	$\rho(1)$	$\rho(2)$	$\rho(3)$	$\rho(4)$	$\rho(5)$	$\rho(6)$	$\rho(7)$	$\rho(8)$	$\chi^2(8)$
	0.03 (-1.71)*	-0.23 (-1.67)*	-0.12 (-0.82)	-0.17 (-1.32)	-0.29 (-1.97)**	-0.38 (-2.11)**	-0.44 (-2.21)**	-0.54 (-2.44)**	12.41 (0.13)
	-0.20 (-2.59)**	-0.10 (-1.35)	-0.34 (-2.74)**	-0.46 (-3.02)**	-0.58 (-3.26)**	-0.57 (-2.99)**	-0.54 (-2.89)**	-0.60 (-2.95)**	26.09 (0.00)*
Y	0.05 (-1.12)	0.11 (0.34)	0.04 (0.33)	-0.14 (-0.70)	-0.22 (-0.77)	-0.21 (-0.46)	-0.18 (-0.28)	-0.18 (-0.20)	5.24 (0.73)
	-0.12 (-1.42)	-0.23 (-1.15)	-0.31 (-1.05)	-0.30 (1.05)	-0.33 (-1.07)	-0.34 (-0.91)	-0.28 (-0.65)	-0.35 (-0.96)	20.84 (0.01)*
	0.09 (-0.82)	-0.25 (-1.78)*	-0.44 (-2.65)**	-0.48 (-2.84)**	-0.32 (-1.57)	-0.27 (-1.33)	-0.31 (-1.93)*	-0.48 (-4.56)**	20.98 (0.01)*
LDS	-0.08 (-1.76)*	-0.03 (-1.22)	-0.13 (-0.96)	-0.26 (-1.40)	-0.30 (-1.14)	-0.32 (-1.17)	-0.38 (-1.41)	-0.49 (-2.26)**	9.52 (0.30)
	0.10 (-1.03)	-0.19 (-2.03)**	-0.22 (-1.63)	-0.40 (-1.95)*	-0.44 (-2.69)**	-0.47 (-3.60)**	-0.49 (-2.48)**	-0.53 (-2.37)**	10.05 (0.26)
	0.229	0.063	0.008	-0.028	-0.055	-0.079	-0.103	-0.126	

i) represents the first-order autocorrelation for the ith difference of the natural log of the real e rate. Numbers in parentheses are t-statistics for the hypothesis that the estimated coefficient l to the simulated bias. Standard errors are calculated using the covariance matrix estimator ed by Hansen and Hodrick (1981). $\chi_2(8)$ is a joint test that all 8 reported autocorrelations are o the estimated bias and is calculated according to the standard errors suggested by Richardson * (**) indicates significance at the 10% (5%) level.

Table 3A
Variance Ratios (CPI) - monthly ending 1988:12

	q =	2	4	8	16	32
Belgium	Mr	1.010	1.182	1.317	1.913	3.419
	Z1	(0.138)	(1.321)	(1.452)*	(2.752)**	(4.797)**
	Z2	[0.112]	[1.153]	[1.383]	[2.709]**	[4.818]**
Canada		0.934	0.844	0.913	1.157	1.459
		(-0.898)	(-1.129)	(-0.399)	(0.472)	(0.911)
		[-0.767]	[-1.018]	[-0.381]	[0.476]	[0.970]
France		0.982	1.137	1.360	1.809	3.046
		(-0.257)	(0.996)	(1.651)*	(2.439)**	(4.057)**
		[-0.227]	[0.934]	[1.647]*	[2.482]**	[4.115]**
Germany		0.994	1.148	1.254	1.758	2.834
		(-0.077)	(1.070)	(1.165)	(2.284)**	(3.637)**
		[-0.065]	[0.962]	[1.131]	[2.285]*	[3.615]**
Italy		1.042	1.203	1.426	1.953	2.560
		(0.576)	(1.470)*	(1.954)**	(2.871)**	(3.094)**
		[0.530]	[1.326]	[1.831]*	[2.765]**	[3.226]**
Japan		1.098	1.240	1.502	1.935	2.324
		(1.336)	(1.739)*	(2.304)*	(2.818)**	(2.626)**
		[1.211]	[1.635]*	[2.163]*	[2.622]**	[2.548]*
Netherlands		0.993	1.145	1.257	1.738	2.804
		(-0.095)	(1.054)	(1.178)	(2.224)**	(3.578)**
		[-0.076]	[0.915]	[1.116]	[2.185]*	[3.596]**
Switzerland		1.019	1.168	1.307	1.472	2.136
		(0.265)	(1.217)	(1.407)*	(1.423)	(2.253)**
		[0.244]	[1.129]	[1.338]	[1.396]	[2.203]*
UK		1.095	1.133	1.215	1.527	2.205
		(1.298)	(0.962)	(0.987)	(1.587)*	(2.389)**
		[1.334]	[1.018]	[1.006]	[1.594]	[2.471]*
FF/DM		1.091	1.294	1.159	1.515	1.251
		(1.246)	(2.130)**	(0.730)	(1.551)*	(0.497)
		[1.004]	[1.650]*	[0.573]	[1.464]	[0.493]
Yen/SF		0.983	1.061	1.237	1.008	0.515
		(-0.237)	(0.444)	(1.089)	(0.024)	(-0.961)
		[-0.217]	[0.401]	[0.985]	[0.022]	[-0.854]
	nq =	186	184	184	176	160
5% Critical Values						
	Z1	1.747	1.778	1.853	2.063	2.180
	Z2	2.013	2.055	2.146	2.458	2.607

* (**) - significant at the 10% (5%) level

Table 3B
Variance Ratios (WPI) - monthly ending 1988:12

	q =	2	4	8	16	32
Belgium	Mr	0.916	1.012	0.884	1.232	2.105
	Z1	(-1.145)	(0.084)	(-0.531)	(0.699)	(2.191)**
	Z2	[-1.086]	[0.076]	[-0.490]	[0.685]	[2.133]*
Canada		0.900	0.756	0.746	0.852	1.035
		(-1.360)*	(-1.771)**	(-1.165)	(-0.447)	(0.069)
		[-1.213]	[-1.604]*	[-1.113]	[-0.458]	[0.074]
Germany		0.990	1.141	1.181	1.592	2.498
		(-0.137)	(1.021)	(0.828)	(1.786)*	(2.970)**
		[-0.116]	[0.919]	[0.793]	[1.765]*	[2.916]**
Italy		0.978	1.083	1.198	1.686	2.400
		(-0.305)	(0.598)	(0.910)	(2.068)**	(2.778)**
		[-0.273]	[0.549]	[0.865]	[2.016]*	[2.851]**
Japan		0.986	1.042	1.140	1.315	1.568
		(-0.189)	(0.301)	(0.641)	(0.949)	(1.127)
		[-0.177]	[0.286]	[0.605]	[0.893]	[1.080]
Netherlands		0.979	1.131	1.174	1.609	2.593
		(-0.293)	(0.950)	(0.798)	(1.836)*	(3.159)**
		[-0.241]	[0.847]	[0.762]	[1.798]*	[3.119]**
Switzerland		0.989	1.113	1.145	1.287	1.924
		(-0.157)	(0.818)	(0.663)	(0.865)	(1.832)*
		[-0.149]	[0.782]	[0.644]	[0.851]	[1.793]*
UK		1.106	1.170	1.183	1.448	2.057
		(1.446)*	(1.232)	(0.838)	(1.351)	(2.095)*
		[1.559]	[1.331]	[0.862]	[1.379]	[2.175]*
Yen/SF		0.933	0.934	1.007	0.807	0.458
		(-0.917)	(-0.482)	(0.031)	(-0.582)	(-1.075)
		[-0.797]	[-0.416]	[0.027]	[-0.517]	[-0.947]
	nq =	186	184	184	176	160
5% Critical Values						
	Z1	1.747	1.778	1.853	2.063	2.180
	Z2	2.013	2.055	2.146	2.458	2.607

* (**) - significant at the 10% (5%) level

Table 3C
Variance Ratios - monthly - 1979-88

		CPI		WPI	
	q =	8	16	8	16
Belgium	Mr	1.408	2.132	0.951	1.360
	Z1	(1.553)*	(2.899)**	(-0.187)	(0.993)
	Z2	[1.512]	[3.012]**	[-0.180]	[0.959]
Canada		0.797	1.079	0.592	0.710
		(-0.771)	(0.203)	(-1.555)**	(-0.741)
		[-0.787]	[0.213]	[-1.519]*	(-0.763]
France		1.411	1.967		
		(1.567)*	(2.476)**		
		[1.656]	[2.689]**		
Germany		1.406	2.031	1.359	1.882
		(1.545)*	(2.639)**	(1.369)	(2.260)**
		[1.529]	[2.751]**	[1.341]	[2.324]*
Italy		1.592	2.251	1.437	2.006
		(2.256)**	(3.202)**	(1.667)*	(2.576)**
		[2.267]**	[3.317]**	[1.670]*	[2.638]**
Japan		1.482	1.883	1.125	1.313
		(1.837)*	(2.261)**	(0.475)	(0.801)
		[1.866]*	[2.332]*	[0.483]	[0.822]
Netherlands		1.442	1.993	1.405	1.927
		(1.683)*	(2.542)**	(1.542)*	(2.374)**
		[1.595]*	[2.598]**	[1.477]	[2.414]*
Switzerland		1.208	1.515	1.170	1.419
		(0.792)	(1.319)	(0.649)	(1.072)
		[0.785]	[1.362]	[0.646]	[1.103]
UK		1.267	1.592	1.221	1.512
		(1.018)	(1.515)*	(0.841)	(1.312)
		[1.085]	[1.595]	[0.897]	[1.386]
FF/DM		0.935	1.448		
		(-0.248)	(1.148)		
		[-0.321]	[1.268]		
Yen/SF		1.331	1.281	1.077	1.063
		(1.262)	(0.719)	(0.295)	(0.161)
		[1.138]	[0.646]	[0.251]	[0.142]
5% critical values					
	Z1	1.919	2.063		
	Z2	2.258	2.454		

Table 3D

Variance Ratios - Annual Data 1900 - 1987

		q = 2	4	8	16
Canada	Mr	1.057	0.833	0.742	0.427
	Z1	(0.527)**	(-0.818)**	(-0.779)**	(-1.164)**
	Z2	[0.504]**	[-0.736]**	[-0.742]*	[-1.176]**
France		0.857	0.779	0.448	0.195
		(-1.325)**	(-1.082)**	(-1.670)**	(-1.635)**
		[-0.823]**	[-0.682]**	[-1.208]**	[-1.387]**
Germany		1.105	1.180	1.068	0.934
		(0.970)*	(0.884)	(0.204)	(-0.134)
		[0.690]**	[0.614]	[0.149]	[-0.109]
Italy		0.933	0.735	0.561	0.449
		(-0.623)**	(-1.300)**	(-1.329)**	(-1.120)**
		[-0.255]**	[-0.605]**	[-0.750]*	[-0.777]*
Japan		1.133	0.867	0.491	0.321
		(1.230)	(-0.651)**	(-1.540)**	(-1.379)**
		[0.875]*	[-0.411]**	[-1.074]**	[-1.137]**
Netherlands		0.992	0.921	0.703	0.393
		(-0.073)**	(-0.385)**	(-0.898)**	(-1.233)**
		[-0.049]**	[-0.279]**	[-0.704]*	[-1.074]**
United Kingdom		1.144	0.897	0.537	0.318
		(1.334)	(-0.504)**	(-1.401)**	(-1.386)**
		[1.154]*	[-0.456]**	[-1.339]**	[-1.406]**
	nq =	86	84	80	64
5% Critical Values					
	Z1	0.838	-0.138	-0.739	-1.012
	Z2	0.873	-0.140	-0.760	-1.067

* (**) - significant at the 10% (5%) level

Figure 3

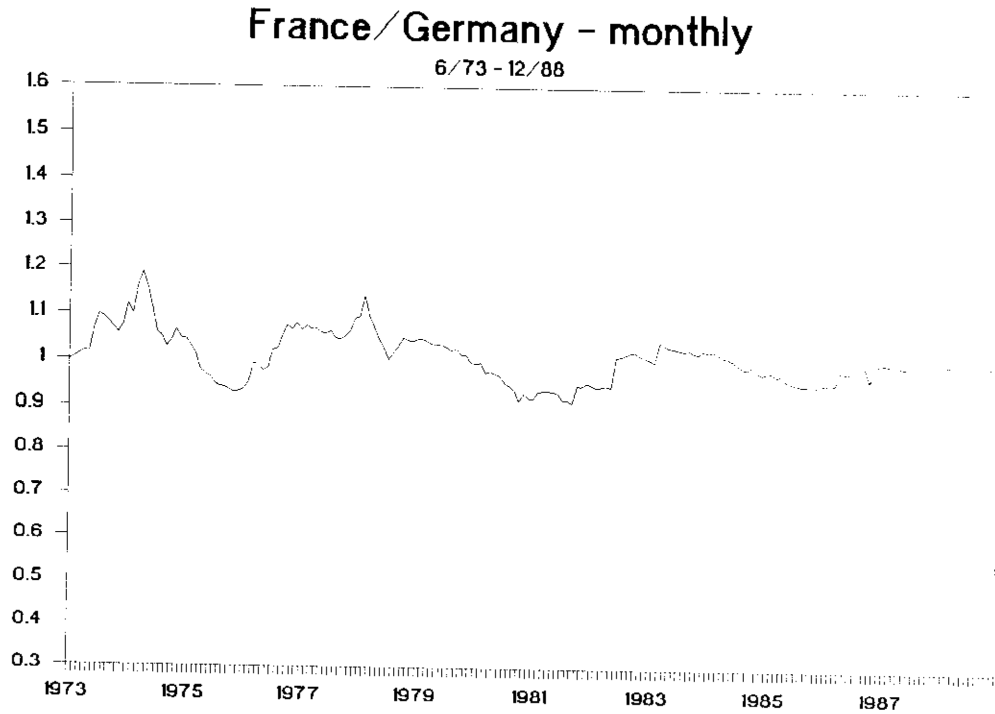

Figure 4

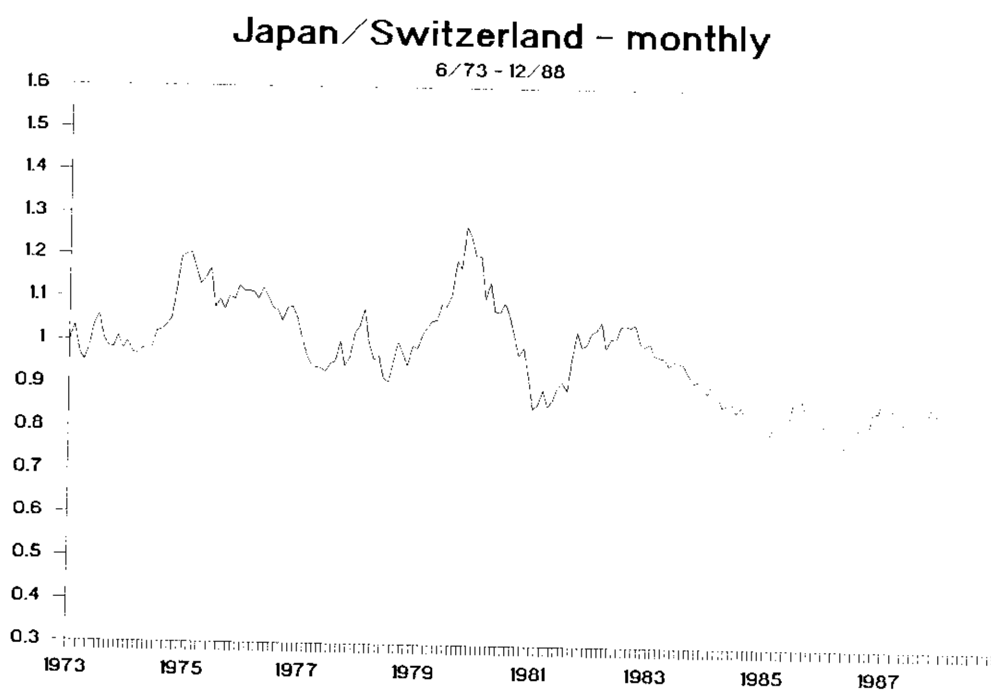

Figure 1

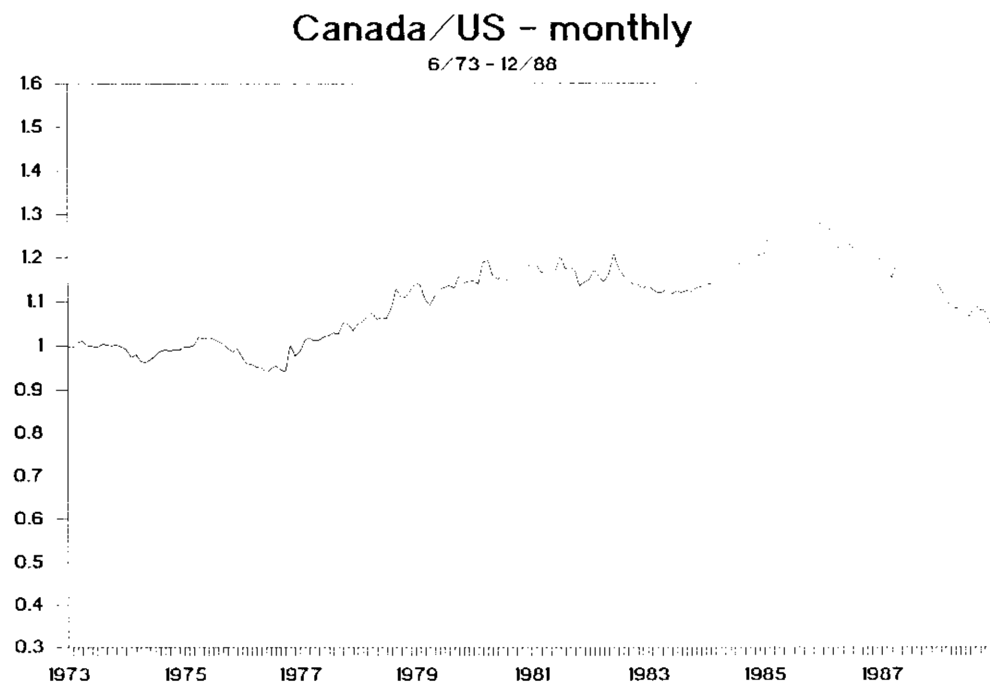

Figure 2